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# Term structure evidence on interest rate smoothing and monetary policy inertia ☆

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**Abstract**

Numerous studies have used quarterly data to estimate monetary policy rules or reaction functions that appear to exhibit a very slow partial adjustment of the policy interest rate. The conventional wisdom asserts that this gradual adjustment reflects a policy inertia or interest rate smoothing behavior by central banks. However, such quarterly monetary policy inertia would imply a large amount of forecastable variation in interest rates at horizons of more than 3 months, which is contradicted by evidence from the term structure of interest rates. The illusion of monetary policy inertia evident in the estimated policy rules likely reflects the persistent shocks that central banks face.


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*Keywords:* Monetary policy; Term structure of interest rates; Taylor rule

## 1. Introduction

How quickly do central banks adjust monetary policy in response to developments in the economy? A common view among economists is that the short-term policy interest rate in many countries is changed at a very sluggish pace over several

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quarters. The evidence supporting this view is found in the many monetary policy rules or reaction functions estimated in the literature with quarterly data. These policy rules take the general partial adjustment form $i_t = (1-\rho)i_t^* + \rho i_{t-1}$, where $i_t$ is the level of the policy interest rate in quarter $t$, which is set as a weighted average of the current desired level, $i_t^*$, and last quarter's actual value, $i_{t-1}$. Based on historical data, estimates of $\rho$ are often in the range of 0.8, so these empirical rules appear to imply a very slow speed of adjustment of the policy rate to its fundamental determinants. This gradual adjustment of the policy rate over several quarters to its desired level is widely interpreted as evidence of an "interest rate smoothing" or "monetary policy inertia" behavior by central banks. For example, Clarida et al. (2000, pp. 157–158) describe their U.S. estimates of various partial adjustment policy rules: "... the estimate of the smoothing parameter $\rho$ is high in all cases, suggesting considerable interest rate inertia: only between 10 and 30 percent of a change in the [desired interest rate] is reflected in the Funds rate within the quarter of the change. Thus, our estimates confirm the conventional wisdom that the Federal Reserve smooths adjustments in the interest rate". Some of the many other recent papers with a similar inertial interpretation of monetary policy rules include Woodford (1999), Goodhart (1999), Levin et al. (1999), Amato and Laubach (1999), and Sack (1998).

Furthermore, a few researchers have also argued recently that the monetary policy inertia apparently present in the real world may be an optimal behavioral response on the part of central banks. For example, one popular such normative argument contends that the quarterly policy inertia and interest rate smoothing behavior helps the central bank focus the expectations of agents in the economy on its stabilization goals and thereby achieve a better outcome (e.g., Levin et al. 1999; Woodford, 1999; Sack and Wieland, 2000).

There is another quite separate literature on "interest rate smoothing", which, at least superficially, may appear to be consistent with the quarterly interest rate smoothing described above. This earlier literature analyzes changes in policy interest rates on a day-by-day basis. Both in the U.S. (e.g., Goodfriend, 1991; Rudebusch, 1995) and in Europe, Japan, and Australia (e.g., Goodhart, 1997; Lowe and Ellis, 1997), central banks appear to follow a pattern of behavior in which changes in the policy rate are undertaken at discrete intervals and in discrete amounts.[1] For example, Rudebusch (1995, p. 264) defines short-term (or weekly) interest rate smoothing as the Fed adjusting interest rates "... in limited amounts... over the course of several weeks with gradual increases or decreases (but not both)...".[2]

[1] Also, see Balduzzi et al. (1997), Dotsey and Otrok (1995), and Eijffinger et al. (1999).

[2] The short-term interest rate smoothing literature distinguishes three interest rates: the market rate at which funds are actually traded, $i_t^{\mathrm{m}}$; the "target" rate that the central bank enforces in the market on a week-by-week basis, $i_t$; and the desired rate, $i_t^*$, that the central bank would set as its target if unconstrained by a desire to adjust the target rate slowly. Note that the "target" rate is not the desired rate. Furthermore, although the market and target rates, which are the ones reported in the popular press, can differ substantially on any given day, they are largely indistinguishable on a monthly average basis as the central bank hits its target, so both are denoted as $i_t$ in this paper (which considers quarterly average data). As examples, Rudebusch (1995) explicitly models $i_t^{\mathrm{m}}$ and $i_t$ on a daily basis (with $i_t^*$ implicit), while Dotsey and Otrok (1995) model $i_t^*$ and $i_t$ on a monthly average basis (so $i_t^{\mathrm{m}} = i_t$).

Many have assumed—including the monetary policy rule papers cited above—that such short-term interest rate smoothing implies the quarterly interest rate smoothing found in the empirical policy rules. However, the earlier short-term interest rate smoothing refers to a partial adjustment over the course of several *weeks*, while quarterly interest rate smoothing refers to a partial adjustment over the course of several *quarters*. With such disparate time frames, the two types of partial adjustment are in fact largely independent, so a central bank could conduct either type of smoothing without much of the other. Indeed, an important point in the short-term interest rate smoothing literature is that although central banks smooth interest rates on a week-to-week or month-to-month basis, there is essentially no quarterly interest rate smoothing. This description follows Mankiw and Miron (1986, p. 225), who note that the postwar term structure of interest rates suggests that at a quarterly frequency "... while the Fed might change the short rate in response to new information, it always (rationally) expected to maintain the short rate at its current level". Goodfriend (1991, p. 10) provides an identical random-walk characterization of the policy rate and argues that changes in the rate set by the Fed "... are essentially unpredictable at forecast horizons longer than a month or two". Similarly, Rudebusch (1995, p. 264) characterizes the Fed's behavior as, "... beyond a horizon of about a month, there are no planned movements to react to information already known". Thus, the earlier short-term interest rate smoothing literature rejects any partial adjustment or policy inertia at a quarterly frequency.[3]

This paper argues that quarterly interest rate smoothing (or monetary policy inertia) is a very modest phenomenon in practice, which accords with the earlier characterization of monetary policy partial adjustment as involving only a very short-term smoothing of rates. This argument, however, must account for the many estimated policy rules that appear to indicate that a high degree of quarterly interest rate smoothing is present in the real world. This seemingly straightforward descriptive evidence of slow adjustment from the inertial empirical policy rules is summarized in the next section, while Section 3 outlines the related normative arguments for the optimality of inertial behavior in a New Keynesian model of output and inflation.

Evidence against the existence of an inertial policy rule is obtained from the behavior of market interest rates at the short-term end of the yield curve. As documented in Section 4, there appears to be very little information generally available in financial markets regarding future interest rate movements beyond the next 1 or 2 months, which is consistent with the results of Mankiw and Miron (1986) and many others. In contrast, Section 5 derives the term structure implications of monetary policy inertia in a New Keynesian model and shows that the large $\rho$ in an inertial rule implies that typically there *are* predictable future changes in the policy rate, which under rational expectations should be embodied in the term structure. Thus, there is an inconsistency between the term structure implications of quarterly

[3] This can be shown formally by simulating the models of short-term interest rate smoothing and fitting a partial adjustment model to the quarterly averaged simulated data. For example, data on the desired and actual funds rates ($i_t^*$ and $i_t$) can be generated according to Eqs. (5) and (6) in Dotsey and Otrok (1995), and after quarterly averaging, the $\hat{\rho}$ in an estimated partial adjustment policy rule is only about 0.1.

interest rate smoothing and the historical term structure evidence. Furthermore, this inconsistency is robust to a variety of different assumptions about the specification of the model and the policy rule.

Assuming financial markets process information efficiently, the term structure evidence implies that the empirical policy rules displaying substantial partial adjustment are misspecified. Section 6 argues that such partial adjustment could be spuriously attributed to a non-inertial central bank, that is, one that displays no quarterly interest rate smoothing. This argument is based on the econometric near-observational equivalence of the partial adjustment rule and a non-inertial rule with serially correlated shocks. That is, significant persistent shocks may explain the illusion of monetary policy inertia, and the conventional empirical partial adjustment rules are misspecified. Furthermore, when monetary policymakers respond to current information—including the persistent shock—interest rate predictability is quite low, which is consistent with the term structure evidence.

## 2. The policy inertia in estimated rules

Many recent studies have estimated models of central bank behavior. A sizable fraction of these empirical policy rules or reaction functions follow Taylor (1993), who proposed a simple rule for monetary policy that sets the quarterly average level of the short-term policy interest rate ($i_t$) in response to (four-quarter) inflation ($\bar{\pi}_t$) and the output gap ($y_t$):

$$i_t = r^* + \bar{\pi}_t + 0.5(\bar{\pi}_t - \pi^*) + 0.5y_t, \tag{1}$$

where $r^*$ is the equilibrium real rate and $\pi^*$ is the inflation target. However, for an empirical version of this rule with estimated response coefficients, a lagged policy rate is also usually included. Accordingly, a typical rule regression has the generic partial adjustment form (ignoring constants) of

$$i_t = (1 - \rho_1)(g_\pi \bar{\pi}_t + g_y y_t) + \rho_1 i_{t-1} + \xi_t, \tag{2}$$

where $\rho_1$, $g_\pi$, and $g_y$ are the coefficients of what is denoted here as Rule 1.

For example, a least-squares regression of Rule 1 on U.S. data from 1987:Q4 to 1999:Q4 yields (ignoring constants)

$$\begin{aligned} i_t = 0.27 \quad & \underset{(0.30)}{(1.53\bar{\pi}_t} \quad + \quad \underset{(0.14)}{0.93y_t)} \quad + \quad \underset{(0.07)}{0.73i_{t-1}} + \xi_t, \\ \sigma_\xi = 0.36, \quad & \bar{R}^2 = 0.96, \end{aligned} \tag{3}$$

where the interest rate is the quarterly average federal funds rate.[4] In this regression, the estimated values of the response coefficients—namely, $g_\pi = 1.53$ for the inflation response and $g_y = 0.93$ for the output response—are just above the 1.5 and 0.5 that

[4] Inflation is defined using the GDP chain-weighted price index (denoted $P_t$, so $\pi_t = 400(\ln P_t - \ln P_{t-1})$ and $\bar{\pi}_t = \frac{1}{4}\sum_{j=0}^{3} \pi_{t-j}$), and the output gap is defined as the percent difference between actual real GDP ($Q_t$) and potential output ($Q_t^*$) estimated by the Congressional Budget Office (i.e., $y_t = 100(Q_t - Q_t^*)/Q_t^*$).

Taylor (1993) originally proposed. Similar estimates are obtained in other empirical studies.[5] Most notable, however, is the large and highly significant estimates of the coefficient on the lagged policy rate, $\hat{\rho}_1 = 0.73$. Indeed, such significant lagged dependence in the empirical estimation of Rule 1 is an extremely robust result in the literature. For example, across six different quarterly U.S. data samples (differing in output gap definitions), Kozicki (1999) reports a range of $\hat{\rho}_1$ from 0.75 to 0.82, while across 16 different quarterly samples of U.S. data (differing in output gap, inflation, and sample period definitions), Amato and Laubach (1999) report a range of $\hat{\rho}_1$ from 0.78 to 0.92.

In contrast to Eq. (3), the regression of the non-inertial form of Rule 1, which imposes the constraint that $\rho_1 = 0$, yields

$$\begin{aligned} i_t = \quad & \underset{(0.13)}{1.59\bar{\pi}_t} \quad + \quad \underset{(0.09)}{0.68 y_t} + \xi_t, \\ \sigma_\xi = 0.73, \quad & \bar{R}^2 = 0.84, \quad DW = 0.33, \end{aligned} \tag{4}$$

which has a significantly worse fit and severely serially correlated errors, although the estimates of $g_\pi$ and $g_y$ are not very different.[6]

The evidence for significant lagged dependence is also robust across different variations of the Taylor rule. In particular, Clarida et al. (2000) recommend a forecast-based specification of the Taylor rule, which I denote as Rule 2,

$$i_t = (1 - \rho_2)(g_\pi \mathrm{E}_{t-1}\bar{\pi}_{t+4} + g_y \mathrm{E}_{t-1} y_t) + \rho_2 i_{t-1} + \xi_t, \tag{5}$$

where $\mathrm{E}_{t-1}\bar{\pi}_{t+4}$ is the forecast of annual inflation five quarters ahead based on the $t-1$ information set and $\mathrm{E}_{t-1} y_t$ is the forecast of the time $t$ output gap based on the $t-1$ information set. An instrumental variables estimate of Rule 2 over the 1987–1999 sample is[7]

$$\begin{aligned} i_t = 0.21 \quad & (\underset{(0.50)}{1.40\mathrm{E}_{t-1}\bar{\pi}_{t+4}} \quad + \quad \underset{(0.28)}{0.90\mathrm{E}_{t-1} y_t}) \quad + \quad \underset{(0.06)}{0.79 i_{t-1}} + \xi_t, \\ \sigma_\xi = 0.41, \quad & \bar{R}^2 = 0.95. \end{aligned} \tag{6}$$

These parameter estimates are broadly similar to ones for this specification given in Clarida et al. (2000, Table 5), although they report even slower partial adjustment with $\hat{\rho}_2 = 0.91$. As above, there is a significant contrast in fit with the estimated non-inertial Rule 2, which has the restriction that $\rho_2 = 0$, although again the sizes of the

[5] See, for example, Kozicki (1999), Amato and Laubach (1999), Sack (1998), and Judd and Rudebusch (1998).

[6] In all the regressions in this section, robust standard errors for the coefficients are reported in parentheses. For example, there is some residual serial correlation in Eq. (3) as well, but for simplicity this paper just considers first-order autoregressive terms.

[7] Four lags of inflation, the output gap, and the interest rate are used as instruments.

$\hat{g}_\pi$ and $\hat{g}_y$ are similar,

$$\begin{aligned} i_t = &\ \underset{(0.32)}{1.33}\mathrm{E}_{t-1}\bar{\pi}_{t+4} + \underset{(0.18)}{0.59}\mathrm{E}_{t-1}y_t + \xi_t, \\ &\sigma_\xi = 1.09, \quad \bar{R}^2 = 0.65, \quad DW = 0.35. \end{aligned} \tag{7}$$

In short, as many have noted, the partial adjustment forms of Rules 1 and 2 appear to fit the data significantly better than those without partial adjustment. This significant lagged dependence in empirical Taylor-type rules also appears to be a quite general feature found in a variety of countries in Europe and elsewhere. For example, Clarida et al. (1998) estimate Rule 2 on quarterly European data and obtain estimates of $\rho_2$ above 0.90 in Germany, France, Italy, and the U.K., and Nelson (2000) provides estimates of Rule 1 for the U.K. that also display significant lagged dependence.

Along with the uniform results on the size and significance of the lagged policy interest rate in the empirical rules, there is also a standard partial adjustment interpretation of this term. When $\rho_1$ or $\rho_2$ equals zero, the current policy rate is based solely on *current* macroeconomic performance (actual or expected). When these lag coefficients are positive (but less than one), then the current policy rate is set equal to a weighted average of this current desired interest rate and last quarter's rate. This conventional wisdom of quarterly monetary policy partial adjustment has been advanced by numerous authors, including Goodhart (1999), Levin et al. (1999), Woodford (1999), Amato and Laubach (1999), Clarida et al. (2000), and Sack and Wieland (2000). Such partial adjustment behavior is typically termed "interest rate smoothing" because the resulting interest rate series will be less volatile than would be suggested by the determinants of policy. Indeed, the degree of quarterly interest rate smoothing or inertia is often *measured* by the size of the speed of adjustment coefficient because as $\rho_1$ or $\rho_2$ increases for a given policy rule, the standard deviation of $\Delta i_t$ falls.[8]

Given the simple forms of Rules 1 and 2, it may seem that the significant estimated partial adjustment could reflect a misspecified reaction function. One such misspecification might involve structural shifts in the parameters of the policy rule for different policy regimes, which might account for the significant lagged interest rate even if the central bank was non-inertial; however, Clarida et al. (2000) and others provide rule estimates over numerous subsamples, and all display a large partial adjustment lag coefficient. Alternatively, Rules 1 and 2 may be misspecified because of the omission of a persistent, serially correlated variable that influences monetary policy. Such an omitted variable could also produce the spurious appearance of partial adjustment in the estimated rule. However, in a wide variety of less parsimonious specifications of $i_t^*$, significant estimates of $\rho$ are still obtained. For example, McNees (1992), McCallum and Nelson (1999), and Fair (2000) estimate

[8] This is true for a single stochastic equation but is, of course, not necessarily true in the context of a complete model. For example, in the Rudebusch and Svensson (1999) model, increasing $\rho_1$ can increase the variance of $\Delta i_t$ and even lead to dynamic instability. Also, a rule with a larger autoregressive coefficient than another rule does not necessarily produce smoother interest rates, because the policy rate volatility also depends on the volatility of the other arguments of the rules.

more complicated structural monetary policy rules and obtain significant evidence of policy inertia with partial adjustment coefficients on the order of 0.8 or higher. In addition, numerous monetary VAR estimates, which provide a very popular implementation of an empirical policy reaction function, also show significant inertia. The estimated VAR interest rate equations contain large and significant lagged interest rate coefficients despite including a wide variety of other regressors. For example, Rudebusch (1998, Table 2) reports the sum of the lagged funds rate coefficients in the reduced form of a well-known quarterly VAR interest rate equation (which has 24 non-interest-rate regressors) as 0.95.[9]

Still, this paper takes issue with the "conventional wisdom" that quarterly monetary policy inertia exists and argues that the common empirical monetary policy rules are indeed misspecified. However, as described below, this misspecification appears difficult to detect directly; thus, this paper focuses on indirect term structure evidence of the misspecification. As a first step, the next section introduces a model of the economy and considers the optimality of policy inertia.

## 3. Optimal monetary policy inertia

The above empirical policy rules imply a very slow speed of adjustment. A $\hat{\rho}_1$ or $\hat{\rho}_2$ of 0.8 implies a 20 percent adjustment *each quarter*, so in a year, a central bank would complete only 60 basis points of a desired one percentage point change. Still, such sluggish behavior may be optimal for a central bank. An obvious explanation is that $i_{t-1}$ is likely an important state variable, so the fully optimal instrument rule would include a response to its value (e.g., Rudebusch and Svensson, 1999).[10] An important example of this occurs in an explicitly forward-looking model, where partial adjustment can be optimal if the private sector is forward looking and the monetary policymaker is credibly committed to a gradual policy rule (see Woodford, 1999; Rotemberg and Woodford, 1999; Levin et al., 1999; Sack and Wieland, 2000; Amato and Laubach, 1999). In such a situation, the small inertial changes in the policy interest rate that are expected in the future can have a large effect on current supply and demand and can help the central bank control macroeconomic fluctuations.[11]

This argument can be elucidated within an empirical New Keynesian model. The key aggregate relationships of the simple theoretical version of this model are

$$\pi_t = \mu_\pi E_t\pi_{t+1} + (1-\mu_\pi)\pi_{t-1} + \alpha_y y_t + \varepsilon_t, \tag{8}$$

$$y_t = \mu_y E_t y_{t+1} + (1-\mu_y)y_{t-1} - \beta_r(i_t - E_t\pi_{t+1} - r^*) + \eta_t, \tag{9}$$

[9] Rudebusch (1998) and Goodfriend (2000) criticize the monetary policy partial adjustment in recent VARs as implausible.

[10] As noted by Levin et al. (1999), this may be especially true for restricted rules, such as Rules 1 and 2, because the lagged policy rate may proxy for excluded lags of other variables.

[11] As a second reason why partial adjustment may be optimal, Sack (2000), Sack and Wieland (2000), and Söderström (2000) cite multiplicative parameter uncertainty; however, the results of Rudebusch (2001) and Peersman and Smets (1999) indicate that the effect of such uncertainty is quite modest empirically.

where $\mathrm{E}_t\pi_{t+1}$ and $\mathrm{E}_t y_{t+1}$ are the expectations of period $t+1$ inflation and output conditional on a time $t$ information set. Much of the appeal of this model lies in its foundations in a dynamic general equilibrium model with nominal price rigidities.[12] An empirical version of this model suitable for quarterly data, where longer leads and lags appear appropriate given the institutional length of contracts and delays in information flows and processing, reformulates Eqs. (8) and (9) as

$$\pi_t = \mu_\pi \mathrm{E}_{t-1}\bar{\pi}_{t+3} + (1-\mu_\pi)\sum_{j=1}^{4}\alpha_{\pi j}\pi_{t-j} + \alpha_y y_{t-1} + \varepsilon_t, \tag{10}$$

$$y_t = \mu_y \mathrm{E}_{t-1}y_{t+1} + (1-\mu_y)\sum_{j=1}^{2}\beta_{yj}y_{t-j} - \beta_r(r_{t-1} - r^*) + \eta_t, \tag{11}$$

where $\mathrm{E}_{t-1}\bar{\pi}_{t+3}$ represents the expectation of average inflation over the next year and $r_{t-1}$ is the real rate relevant for output. In particular, $r_{t-1}$ is defined as a weighted combination of an ex ante 1-year rate and an ex post 1-year rate:

$$r_{t-1} = \mu_r(\mathrm{E}_{t-1}\bar{\imath}_{t+3} - \mathrm{E}_{t-1}\bar{\pi}_{t+4}) + (1-\mu_r)(\bar{\imath}_{t-1} - \bar{\pi}_{t-1}), \tag{12}$$

where $\bar{\imath}_t$ is a four-quarter average of past interest rates, i.e., $\bar{\imath}_t = \frac{1}{4}\sum_{j=0}^{3} i_{t-j}$.

This model allows the analysis below to consider a wide range of explicit forward-looking behavior, which is important given the uncertainty about the quantitative importance of expectations. As a theoretical matter, the values of $\mu_\pi$, $\mu_y$, and $\mu_r$ are not clearly determined.[13] Furthermore, the empirical evidence on the appropriate values of these parameters, which is surveyed in Rudebusch (2002), is not decisive. At one extreme, the model with $\mu_\pi$, $\mu_y$, and $\mu_r$ set equal to zero matches the completely adaptive expectations model of Rudebusch and Svensson (1999) and Rudebusch (2001), which has had some success in approximating the time series data in the manner of a small estimated VAR (see Fuhrer, 1997; Estrella and Fuhrer, 1998). In this extreme model, inflation and output are not based on explicit expectations but are based completely on lags (which may implicitly represent adaptive expectations), and the real rate is an average of the past four quarters of real rates (which may represent planning and production lags from interest rates to output or an adaptive expectations version of the term structure as in Modigliani and Schiller, 1973). However, estimated forward-looking models also have had some success in fitting the data, as in Rotemberg and Woodford (1999), Fuhrer (2000), McCallum and Nelson (1999), and Fuhrer and Moore (1995). The analysis below takes a very eclectic view and conditions on a wide range of possible values for $\mu_\pi$, $\mu_y$, and $\mu_r$. In contrast, there is less contention regarding the values of the other

[12] For explicit derivations and discussion, see Woodford (1996), Goodfriend and King (1997), Walsh (1998), Clarida et al. (1999), Svensson (1999a, b), McCallum and Nelson (1999), and Rudebusch (2002).

[13] From well-known contracting models of price-setting behavior, it is possible to derive an inflation equation with $\mu_\pi \approx 1$ (e.g., Roberts, 1995). However, many authors assume that with realistic costs of adjustment and overlapping price and wage contracts there will be some inertia in inflation, so $\mu_\pi$ will be less than one, and with even higher costs for adjusting output, $\mu_y$ is likely much less than one as well. See Svensson (1999a, b) and Fuhrer and Moore (1995) and Fuhrer (1997).

Table 1
Model parameter values

| Coefficient | Value |
|---|---|
| $\alpha_{\pi 1}$ | 0.67 |
| $\alpha_{\pi 2}$ | −0.14 |
| $\alpha_{\pi 3}$ | 0.40 |
| $\alpha_{\pi 4}$ | 0.07 |
| $\alpha_y$ | 0.13 |
| $\beta_{y1}$ | 1.15 |
| $\beta_{y2}$ | −0.27 |
| $\beta_r$ | 0.09 |
| $\sigma_\varepsilon$ | 1.012 |
| $\sigma_\eta$ | 0.833 |

parameters in the model, and these are set equal to the values given in Table 1, which are obtained from the data in Rudebusch (2002) for a very similar model.[14]

In order to calculate optimal monetary policy, I assume a standard loss function in which the central bank minimizes variation in inflation around its target $\pi^*$, the output gap, and changes in the interest rate (see Rudebusch and Svensson, 1999; Clarida et al., 1999):

$$E[L_t] = \mathrm{Var}[\bar{\pi}_t - \pi^*] + \lambda\,\mathrm{Var}[y_t] + v\,\mathrm{Var}[\Delta i_t], \tag{13}$$

where $\Delta i_t = i_t - i_{t-1}$, and the parameters $\lambda \geqslant 0$ and $v \geqslant 0$ are the relative weights on output and interest rate stabilization, respectively, with respect to inflation stabilization. (Note this loss function is only used in this section for the discussion of optimal inertia.)

Table 2 summarizes the optimal amount of monetary policy inertia for various models, rules, and loss functions. The table displays the lag coefficients $\rho_1$ and $\rho_2$ from the *optimal* versions of Rules 1 and 2, across models with a range of forward-looking behavior. In particular, for inflation, $\mu_\pi$ is set equal to 0.1, 0.3, or 0.5 because the many available empirical estimates described in Rudebusch (2002) suggest that a very broad plausible range for $\mu_\pi$ is between 0 and 0.6.[15] Similarly for output, $\mu_y$ is set equal to either 0 or 0.3. Almost all empirical estimates have assumed that $\mu_y = 0$ (e.g., Fuhrer and Moore, 1995), and Fuhrer and Rudebusch (2002) provide empirical support for this value; however, Fuhrer (2000) models a habit persistence model, which suggests that $\mu_y$ is approximately equal to 0.3 (see Rudebusch, 2002). Finally for interest rates, $\mu_r$ is varied over essentially the entire range, so $\mu_r = 0.1$, 0.5, or 0.9, because the multicollinearity of many interest rates makes it hard to obtain decisive empirical evidence on its value (e.g., Fuhrer and Moore, 1995). The coefficients of Rules 1 and 2 are optimized in various models according to two different

[14] These are also little different from the values given in Rudebusch and Svensson (1999); in any case, the qualitative results below are robust to their variation. The estimated constants are not reported.

[15] For example, Fuhrer (1997) estimates $\mu_\pi$ to be about zero, while Fuhrer and Moore (1995) assume $\mu_\pi$ is 0.5.

Table 2
Optimal lag coefficients for policy Rules 1 and 2

| Model | | | Loss: $\lambda = 1$; $v = 0.5$ | | Loss: $\lambda = 1$; $v = 0.1$ | |
|---|---|---|---|---|---|---|
| $\mu_r$ | $\mu_\pi$ | $\mu_y$ | $\rho_1$ | $\rho_2$ | $\rho_1$ | $\rho_2$ |
| 0.1 | 0.1 | 0 | 0.27 | 0.47 | 0.00 | 0.22 |
| 0.1 | 0.3 | 0 | 0.25 | 0.47 | 0.00 | 0.25 |
| 0.1 | 0.5 | 0 | 0.25 | 0.55 | 0.00 | 0.33 |
| 0.1 | 0.1 | 0.3 | 0.16 | 0.40 | 0.00 | 0.14 |
| 0.1 | 0.3 | 0.3 | 0.16 | 0.43 | 0.00 | 0.18 |
| 0.1 | 0.5 | 0.3 | 0.18 | 0.53 | 0.00 | 0.27 |
| 0.5 | 0.1 | 0 | 0.57 | 0.67 | 0.50 | 0.51 |
| 0.5 | 0.3 | 0 | 0.55 | 0.67 | 0.49 | 0.53 |
| 0.5 | 0.5 | 0 | 0.57 | 0.73 | 0.50 | 0.62 |
| 0.5 | 0.1 | 0.3 | 0.56 | 0.69 | 0.50 | 0.56 |
| 0.5 | 0.3 | 0.3 | 0.56 | 0.71 | 0.49 | 0.55 |
| 0.5 | 0.5 | 0.3 | 0.59 | 0.73 | 0.51 | 0.59 |
| 0.9 | 0.1 | 0 | 0.70 | 0.73 | 0.62 | 0.55 |
| 0.9 | 0.3 | 0 | 0.72 | 0.73 | 0.65 | 0.58 |
| 0.9 | 0.5 | 0 | 0.74 | 0.79 | 0.68 | 0.70 |
| 0.9 | 0.1 | 0.3 | 0.74 | 0.75 | 0.67 | 0.58 |
| 0.9 | 0.3 | 0.3 | 0.75 | 0.74 | 0.70 | 0.57 |
| 0.9 | 0.5 | 0.3 | 0.77 | 0.77 | 0.75 | 0.64 |

*Notes*: The optimal lag coefficients for Rules 1 and 2—$\rho_1$ and $\rho_2$, respectively—are reported for each of 18 parameterizations of the model, which have various $\mu_\pi$, $\mu_y$, and $\mu_r$ weights on expectational terms, and for both parameterizations of the loss function. Both loss function parameterizations have equal weight on output and inflation volatility ($\lambda = 1$) but a stronger ($v = 0.5$) or weaker ($v = 0.1$) interest rate smoothing motive. The associated optimal $g_\pi$ and $g_y$ are not reported.

parameterizations of the loss function. Columns 4 and 5 of Table 2 provide the optimal $\rho_1$ and $\rho_2$ with $\lambda = 1$ and $v = 0.5$, the baseline case in Rudebusch and Svensson (1999), while columns 6 and 7 provide the optimal $\rho_1$ and $\rho_2$ with $\lambda = 1$ and $v = 0.1$, which incorporates a very modest incentive to reduce interest rate volatility.[16] With $v$ equal to 0.5 or 0.1, respectively, these loss functions equally penalize a 1 percent output gap, a 1 percentage point inflation gap, and a 1.41 or a 3.16 percentage point quarterly change in the funds rate. This appears to be a plausible range of penalty on interest rate volatility given the various reasons to reduce such volatility that have been proposed in the literature.[17]

[16] The results in Table 2 are obtained by numerically minimizing the loss function over the parameters $g_\pi$, $g_y$, and $\rho_j$ in the model of (10), (11), and Rule $j$. The policy rule is subject to an i.i.d. error with $\sigma_\xi = 0.4$, which is in the range of the empirical estimates in Section 2. As usual, the policy rule is assumed to be perfectly credible, so agents know the rule and assume (correctly) that it will be followed. The results are obtained using the "AIM" algorithm (Anderson and Moore, 1985) available at http://www.federalreserve.gov/pubs/oss/oss4/aimindex.html.

[17] There are three broad such motives (e.g., Lowe and Ellis, 1997). First, interest rate volatility may induce instability in financial markets (e.g., Goodfriend, 1991; Rudebusch, 1995; Cukierman, 1996). Second, large interest rate changes may be difficult to achieve politically because of the decision-making process (e.g., Goodhart, 1997) or because such changes may be taken as an adverse signal of inconsistency

As is evident in Table 2, a large range of optimal lag coefficients—between 0 and 0.8—can be rationalized for some combination of model and loss function. Surprisingly, there is little dependence of the optimal $\rho_1$ or $\rho_2$ on the values of $\mu_\pi$ or $\mu_y$. Instead, the degree of optimal quarterly interest rate smoothing is crucially dependent on the value of $\mu_r$, which determines the degree to which interest rate expectations are forward looking. This is consistent with the interpretation of Woodford (1999) and Levin et al. (1999) that policy inertia is optimal when it alters expectations of future interest rates that are also important determinants of current demand.

## 4. Term structure evidence on interest rate predictability

The preceding two sections documented the large and significant estimated coefficient on the lagged interest rate in quarterly central bank reaction functions as well as the optimality of such monetary policy partial adjustment or inertia when economic agents are forward looking with respect to future interest rate movements. This section focuses on measuring how much financial market participants actually know about future interest rate movements. This evidence will provide some crucial benchmarks for the next section which quantifies the term structure implications of monetary policy inertia.

The partial adjustment of monetary policy by a central bank suggests that there are forecastable future movements in the policy interest rate. The amount of such forecastable variation can be measured with a standard term structure regression such as

$$i_{t+j} - i_{t+j-1} = \delta + \gamma(\mathrm{E}_t i_{t+j} - \mathrm{E}_t i_{t+j-1}) + \psi^j_{t+j} \tag{14}$$

(for $j \geqslant 1$). This equation regresses the realized change in the policy rate between two adjacent quarters on the expected such change.[18] Under rational expectations, $i_{t+j} = \mathrm{E}_t i_{t+j} + e_{t+j}$, where $e_{t+j}$, the expectational error, has a mean of zero and is uncorrelated with time $t$ information. In this case, the interest rate forecasting regression equation (14) would yield in the limit an estimate of $\hat{\delta} = 0$ and $\hat{\gamma} = 1$. Furthermore, for assessing monetary policy inertia, a statistic of particular interest is the $R^2$ of this regression, which provides a measure of the forecastability of future interest rate changes (for discussion of such measures, see Diebold and Kilian, 2001).

Many papers have estimated term structure regressions such as Eq. (14) for the postwar period using 3- and 6- or 6- and 12-month Treasury bill spreads as proxies for expectations (see, for example, Mankiw and Miron, 1986; Mishkin, 1988; Cook and Hahn, 1990; Rudebusch, 1995). These studies typically have obtained $R^2$'s very

(*footnote continued*)
and incompetence (e.g., Goodhart, 1999). Finally, smaller interest rate changes seem to make it less likely that the zero bound on nominal interest rates would be reached (though Woodford (1999) disagrees).

[18] These "marginal" regressions are common in the literature (e.g., Mishkin, 1988); however, I obtained similar results with other forms as well.

close to zero. For example, in a 1959–1979 sample, Mankiw and Miron (1986, Table 1) obtain an $R^2$ of 0.02 in a regression of the change in the 3-month rate on the 3- and 6-month spread. However, these results may be too pessimistic because they typically cover a long sample that is unlikely to be a consistent monetary policy regime (see Fuhrer, 1996). In contrast, the term structure implications derived in the next section assume that agents know the policy rule that the central bank is committed to. As a complement to the earlier results, I estimate Eq. (14) with rates on 3-month eurodollar and eurodollar futures, which have been the trading vehicle of choice for hedging short-run future interest rate movements since the mid-1980s.[19] The eurodollar regressions use a short sample from 1988:01 to 2000:01, covering what is arguably a single consistent policy regime.

Denote $ED(t+j)_t$ as the interest rate on eurodollar deposits during quarter $t+j$ that is expected at the end of quarter $t$. Thus, $ED(t+1)_t$ is the spot 3-month eurodollar rate at the end of quarter $t$, and $ED(t+2)_t$ is the rate on a eurodollar futures contract that settles 3 months ahead.[20] Then assume that $ED(t+j)_t = \mathrm{E}_t i_{t+j} + \phi_t^j$, where $\phi_t^j$ is the term premium associated with the $j$th contract. Under the expectations hypothesis of the term structure, the term premia are assumed to be constant over time, but in practice it is widely recognized that there is some time variation. The consequences of time-varying term premia are discussed below.

Using eurodollar data to predict the one-quarter-ahead change in the quarterly average funds rate from 1988:Q1 to 2000:Q1, Eq. (14) with $j = 1$ is estimated as

$$\begin{aligned} i_{t+1} - i_t = \ & \underset{(0.05)}{-0.25} + \underset{(0.10)}{0.83}(ED(t+1)_t - i_t) + \psi_{t+1}^1, \\ \sigma_{\psi 1} = 0.30, \quad & R^2 = 0.57. \end{aligned} \tag{15}$$

This equation indicates that the 3-month eurodollar rate forecasts $\Delta i_{t+1}$ quite well (with an average term premium of about 25 basis points). The $R^2$ indicates that over 50 percent of the one-quarter-ahead variation in the funds rate is known by the end of the preceding quarter. This predictability is consistent with the evidence and interpretation in Rudebusch (1995) of interest rate smoothing at a weekly and monthly frequency. That is, at the end of quarter $t$, financial markets have some information about changes during the first several weeks of the following quarter.[21] In addition, in this regression, changes in the funds rate *during* quarter $t$ (which are of course known at the end of quarter $t$) will also help predict the quarterly average change $\Delta i_{t+1}$. Still, after replacing $i_t$ with the end of quarter $t$ funds rate, substantial predictive power remains with $R^2 \approx 0.3$.

[19] Eurodollar futures contracts are based on the 90-day London Interbank Offered Rate (LIBOR). For further details, see Jegadeesh and Pennacchi (1996).

[20] For the three regressions using eurodollar rates, quarters are defined to start at the eurodollar futures contract settlement dates (which occur about 2 weeks before the start dates of the usual quarters) in order to capture true two- and three-quarter-ahead expectations.

[21] In particular, this significant predictive ability for $\Delta i_{t+1}$ is consistent with the documented ability of a 2- and 1-month interest rate spread to predict the 1-month-ahead change.

Of particular interest in assessing quarterly monetary policy inertia will be the predictive ability at slightly longer horizons. Predicting the one-quarter change in the funds rate two quarters ahead (Eq. (14) with $j = 2$) yields[22]

$$\begin{aligned} i_{t+2} - i_{t+1} = &\underset{(0.07)}{-0.04} + \underset{(0.18)}{0.45}(ED(t+2)_t - ED(t+1)_t) + \psi^2_{t+2}, \\ \sigma_{\psi 2} = 0.42,\quad &R^2 = 0.11. \end{aligned} \tag{16}$$

Predicting $\Delta i_{t+3}$ at quarter $t$ yields

$$\begin{aligned} i_{t+3} - i_{t+2} = &\underset{(0.08)}{-0.06} + \underset{(0.30)}{0.35}(ED(t+3)_t - ED(t+2)_t) + \psi^3_{t+3}, \\ \sigma_{\psi 3} = 0.44,\quad &R^2 = 0.03. \end{aligned} \tag{17}$$

These regressions indicate that there is little if any information usually available in financial markets for predicting the change in the funds rate 3–6 months out ($R^2 = 0.11$) and no information for predicting it 6–9 months out ($R^2 = 0.03$).[23] These $R^2$'s will be used as benchmarks for assessing the plausibility of monetary policy inertia in the next section. These results turn out to be only marginally better than the standard ones described above.[24] The lack of information in these regressions is also consistent with the evidence in Kuttner (2001), where a surprise change in the policy rate target on a particular day shifts the level of the term structure by a similar amount across all horizons, but carries little information about future changes in rates.

Finally, the presence of time-varying term premia should be considered, which, as stressed by Mankiw and Miron (1986), can have important consequences for empirical regressions like Eq. (16). The sample estimates of the $\gamma$ and $R^2$ of this regression will depend positively on the covariance between the independent and dependent variables, $\Delta i_{t+2}$ and $ED(t+2)_t - ED(t+1)_t$, and inversely on their variances. Accordingly, as the time variation in the term premia becomes more significant (boosting the independent, noisy variation in the eurodollar spread), the estimates $\gamma$ and $R^2$ can be driven away from 1 even in the limit. The standard deviation of the residual to the term structure regression provides a rough upper bound on the size of the term premium. For example, in Eq. (16), $\psi^2_t = \phi^1_t - \phi^2_t + e_{t+2} - e_{t+1}$, which is a combination of term premia and the expectational errors. The expectational errors are orthogonal to the term premia; thus, the standard deviation of the term premium associated with the $t+2$ and $t+1$ eurodollar spread (i.e., $\phi^1_t - \phi^2_t$) is smaller than 0.42, the standard deviation of the regression ($\sigma_{\psi 2}$).

[22] For $j > 1$, the forecast errors will have an $MA(j-1)$ moving average correlation, so robust standard errors are reported in parentheses.

[23] For predicting $\Delta i_{t+2}$, the $p$-value for the hypothesis that $R^2 = 0$ is 0.01, and for predicting $\Delta i_{t+3}$, the $p$-value is 0.15.

[24] This slightly better performance may be a spurious small-sample result, perhaps reflecting the unusual 1994 episode discussed below. Also, see Lange et al. (2001).

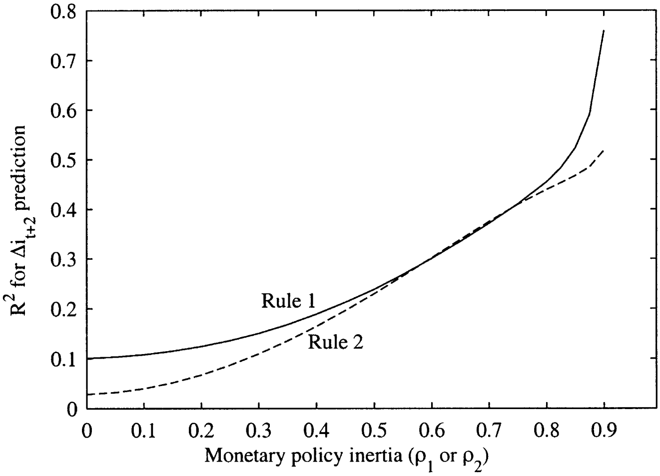


Fig. 1. Interest rate forecastability and policy inertia. This figure shows the direct relationship between the predictability of future changes in interest rates (measured by the $R^2$ of a term structure regression) and the amount of inertia in the policy rule (measured by the lag coefficient $\rho_1$ or $\rho_2$).

## 5. Term structure implications of policy inertia

The previous section provided evidence that beyond a horizon of 3 months there is little predictive information in financial markets about the future path of short-term interest rates. This section explores whether that evidence can be reconciled with a significant degree of quarterly monetary policy inertia. Intuitively, such a reconciliation seems unlikely, for if the funds rate is typically adjusted by only 20 percent toward its desired target in a given quarter, then the remaining 80 percent adjustment should be expected to occur in future quarters. The partial adjustment of the short-term policy interest rate embodied in Rule 1 or 2 with high $\rho_1$ or $\rho_2$ implies that there typically is a large amount of predictable future variation in the policy rate. Indeed, this is the essence of *optimal* policy inertia: Because private agents know that the policy rate is likely to be adjusted by a certain amount in the future, they change their behavior today.

The relationship between the forecastable variation in the interest rate, as measured by the $R^2$ of the $\Delta i_{t+2}$ prediction equation, and quarterly policy inertia, as measured by the $\rho_1$ and $\rho_2$ in Rules 1 and 2, is illustrated in Fig. 1. This figure graphs the (analytical population) value of the $R^2$ of regression equation (14), with $j = 2$, as a function of the value of $\rho_1$ or $\rho_2$ for a representative case of the model described above, namely, with $\mu_\pi = 0.3$, $\mu_r = 0.5$, and $\mu_y = 0$ (and the other parameters given in Table 1).[25] Also, for both policy Rules 1 and 2 (Eqs. (2) and (5)), $g_\pi$ and $g_y$ are set

[25] As above for Table 2, the unique stationary rational expectations solution for each specified policy rule and model is solved via AIM (see Levin et al., 1999; Anderson and Moore, 1985). The reduced-form representation of the saddle-point solution is computed, the unconditional variance–covariance matrix of

Table 3
Predicting $\Delta i_{t+2}$ with various models and rules

| Model | | | Rule 1 | | | | Rule 2 | | | |
|---|---|---|---|---|---|---|---|---|---|---|
| | | | $\rho_1 = 0$ | | $\rho_1 = 0.8$ | | $\rho_2 = 0$ | | $\rho_2 = 0.8$ | |
| $\mu_r$ | $\mu_\pi$ | $\mu_y$ | $R^2_L$ | $R^2_U$ | $R^2_L$ | $R^2_U$ | $R^2_L$ | $R^2_U$ | $R^2_L$ | $R^2_U$ |
| 0.1 | 0.1 | 0 | 0.02 | 0.22 | 0.12 | 0.55 | 0.00 | 0.13 | 0.18 | 0.59 |
| 0.1 | 0.3 | 0 | 0.03 | 0.25 | 0.19 | 0.71 | 0.00 | 0.13 | 0.24 | 0.66 |
| 0.1 | 0.5 | 0 | 0.07 | 0.40 | 0.65 | 0.98 | 0.02 | 0.15 | 0.33 | 0.75 |
| 0.1 | 0.1 | 0.3 | 0.04 | 0.27 | 0.13 | 0.52 | 0.02 | 0.20 | 0.17 | 0.55 |
| 0.1 | 0.3 | 0.3 | 0.04 | 0.28 | 0.19 | 0.65 | 0.03 | 0.20 | 0.20 | 0.59 |
| 0.1 | 0.5 | 0.3 | 0.07 | 0.32 | 0.36 | 0.89 | 0.04 | 0.20 | 0.24 | 0.64 |
| 0.5 | 0.1 | 0 | 0.01 | 0.21 | 0.11 | 0.52 | 0.00 | 0.12 | 0.16 | 0.56 |
| 0.5 | 0.3 | 0 | 0.02 | 0.24 | 0.17 | 0.67 | 0.00 | 0.12 | 0.21 | 0.63 |
| 0.5 | 0.5 | 0 | 0.07 | 0.37 | 0.42 | 0.90 | 0.01 | 0.13 | 0.29 | 0.71 |
| 0.5 | 0.1 | 0.3 | 0.03 | 0.25 | 0.10 | 0.48 | 0.02 | 0.17 | 0.13 | 0.50 |
| 0.5 | 0.3 | 0.3 | 0.03 | 0.25 | 0.14 | 0.58 | 0.02 | 0.17 | 0.16 | 0.54 |
| 0.5 | 0.5 | 0.3 | 0.05 | 0.28 | 0.20 | 0.67 | 0.03 | 0.17 | 0.18 | 0.56 |
| 0.9 | 0.1 | 0 | 0.01 | 0.21 | 0.10 | 0.50 | 0.00 | 0.11 | 0.14 | 0.53 |
| 0.9 | 0.3 | 0 | 0.02 | 0.23 | 0.16 | 0.64 | 0.00 | 0.10 | 0.19 | 0.59 |
| 0.9 | 0.5 | 0 | 0.06 | 0.35 | 0.34 | 0.82 | 0.01 | 0.11 | 0.25 | 0.68 |
| 0.9 | 0.1 | 0.3 | 0.02 | 0.23 | 0.08 | 0.44 | 0.01 | 0.15 | 0.11 | 0.46 |
| 0.9 | 0.3 | 0.3 | 0.02 | 0.23 | 0.11 | 0.50 | 0.01 | 0.15 | 0.12 | 0.48 |
| 0.9 | 0.5 | 0.3 | 0.04 | 0.26 | 0.13 | 0.52 | 0.03 | 0.16 | 0.13 | 0.48 |
| Median | | | 0.03 | 0.25 | 0.15 | 0.61 | 0.01 | 0.15 | 0.18 | 0.57 |

*Notes*: For 18 models, non-inertial ($\rho_1 = 0$) and inertial ($\rho_1 = 0.8$) versions of Rule 1 and non-inertial ($\rho_2 = 0$) and inertial ($\rho_2 = 0.8$) versions of Rule 2 are considered. For each combination of model and rule, 5,000 samples of 100 observations of data are generated. A term structure regression like Eq. (16) is estimated for each sample, and the resulting distributions of $R^2$ values are summarized by $R^2_L$ and $R^2_U$, the 5 percent lower and upper critical values. The median values of these critical values across the 18 models are also reported.

equal to 1.5 and 0.8, respectively, and the rule error is i.i.d. with $\sigma_\xi = 0.4$. (This calibration is in the range of the empirical rule estimates given in Section 2.) Note that even for the non-inertial policy rules there is some predictable future movement in interest rates (with $R^2 = 0.10$ when $\rho_1 = 0$ and $R^2 = 0.03$ when $\rho_2 = 0$). For example, the forecasting power with Rule 1 when $\rho_1 = 0$ reflects the fact that there are predictable changes two quarters ahead in the output gap and in the four-quarter inflation rate, which partly determine future changes in interest rates. Even though the output gap and inflation are highly persistent in levels, the associated slow mean reversion implies only a modest predictability of future quarterly *changes* in these series and in $\Delta i_t^*$. Most importantly, as $\rho_1$ and $\rho_2$ increase, the amount of predictable

(*footnote continued*)
the model variables and the term spreads is obtained analytically, and the term structure regression asymptotic $R^2$ is calculated using the appropriate variances and covariances.

variation in $\Delta i_{t+2}$ also increases, with $R^2$ values of 0.45 at $\rho_1 = 0.8$ and 0.44 at $\rho_2 = 0.8$.[26]

This basic relationship between predictable interest rate variation and monetary inertia is robust across a wide variety of models and rules. Table 3 examines the same 18 different parameterizations of the model considered in Section 3 and non-inertial ($\rho_1 = \rho_2 = 0$) and inertial ($\rho_1 = \rho_2 = 0.8$) versions of both Rules 1 and 2. (Again, $g_\pi = 1.5$, $g_y = 0.8$, and $\sigma_\xi = 0.4$.) In addition, a time-varying term premium, $\phi_t^1 - \phi_t^2$, is included, which is assumed to be i.i.d. with a standard deviation of 0.10.[27] As noted above, such a term premium reduces the $R^2$ values. Each model and rule combination reports $R^2_{\mathrm{L}}$ and $R^2_{\mathrm{U}}$, which are the 5 percent lower and upper critical values, respectively, for the small-sample distribution of the $R^2$ (which are appropriate for 95 percent one-sided or 90 percent two-sided tests). These critical values are calculated from 5,000 simulated samples of the model and the given rule (with 100 observations each), and they allow a probabilistic assessment of the historical term structure regression results given in Section 4. The bottom line in the table gives the median $R^2_{\mathrm{L}}$ and $R^2_{\mathrm{U}}$ values across all models. Given the uncertainty in choosing a single model documented above, I focus on these median values (also, although there is interesting variation across models, the value of $\rho_1$ or $\rho_2$ is the key determinant of interest rate predictability). Based on the historical results with eurodollar data, the benchmark $R^2$ value for the $\Delta i_{t+2}$ prediction regression is 0.11. This value is included in the confidence intervals for the non-inertial $\rho_1 = 0$ and $\rho_2 = 0$ cases; indeed, it is quite close to the small-sample means (which are not shown). In contrast, for the inertial policy rules, the median $R^2$ confidence intervals with $\rho_1 = 0.8$ and with $\rho_2 = 0.8$ both lie above the historical $R^2$ value.

As shown in Table 4, very similar results are obtained for the $\Delta i_{t+3}$ prediction regression. Again, an $R^2_{\mathrm{L}}$ and $R^2_{\mathrm{U}}$ pair is calculated for each of the model and rule combinations used in Table 3. From the previous section, the benchmark value of $R^2$ from the historical data is 0.03. As before, this historical value is contained in the median confidence intervals of the non-inertial policy rules but not in the median inertial policy rule intervals.

In brief, quarterly partial adjustment and interest rate smoothing or inertia do not appear to be consistent with the lack of information in the term structure of interest rates about the future path of interest rates.

[26] The exception to note is that as $\rho_1$ or $\rho_2$ approach one, $i_t$ becomes a random walk and the model is dynamically unstable. However, if Eqs. (2) and (5) were rewritten without the $(1-\rho_i)$ factors, then with $\rho_1$ or $\rho_2$ equal to one, $\Delta i_t$ would take on the persistence properties of the rule arguments.

[27] Mankiw and Miron (1986, Table 3) estimate the standard error of this term premium to be 0.16, while in a more complicated time-series specification with monthly data, Dotsey and Otrok (1995) estimate it to be 0.13. This standard deviation is about one-fourth the size of the regression standard error of Eq. (15), which includes the eurodollar term premia and the orthogonal expectational error. The term premia also reduce the slope estimates in the term structure regression to close to the historical values.

Table 4
Predicting $\Delta i_{t+3}$ with various models and rules

| Model | | | Rule 1 | | | | Rule 2 | | | |
|---|---|---|---|---|---|---|---|---|---|---|
| | | | $\rho_1 = 0$ | | $\rho_1 = 0.8$ | | $\rho_2 = 0$ | | $\rho_2 = 0.8$ | |
| $\mu_r$ | $\mu_\pi$ | $\mu_y$ | $R^2_{\mathrm{L}}$ | $R^2_{\mathrm{U}}$ | $R^2_{\mathrm{L}}$ | $R^2_{\mathrm{U}}$ | $R^2_{\mathrm{L}}$ | $R^2_{\mathrm{U}}$ | $R^2_{\mathrm{L}}$ | $R^2_{\mathrm{U}}$ |
| 0.1 | 0.1 | 0 | 0.01 | 0.21 | 0.05 | 0.47 | 0.00 | 0.13 | 0.06 | 0.46 |
| 0.1 | 0.3 | 0 | 0.02 | 0.22 | 0.11 | 0.65 | 0.00 | 0.13 | 0.09 | 0.53 |
| 0.1 | 0.5 | 0 | 0.06 | 0.37 | 0.59 | 0.98 | 0.01 | 0.14 | 0.16 | 0.64 |
| 0.1 | 0.1 | 0.3 | 0.02 | 0.24 | 0.06 | 0.45 | 0.01 | 0.17 | 0.06 | 0.42 |
| 0.1 | 0.3 | 0.3 | 0.03 | 0.24 | 0.11 | 0.59 | 0.01 | 0.17 | 0.08 | 0.47 |
| 0.1 | 0.5 | 0.3 | 0.06 | 0.30 | 0.29 | 0.87 | 0.02 | 0.18 | 0.11 | 0.52 |
| 0.5 | 0.1 | 0 | 0.01 | 0.20 | 0.04 | 0.44 | 0.00 | 0.11 | 0.05 | 0.43 |
| 0.5 | 0.3 | 0 | 0.01 | 0.22 | 0.10 | 0.61 | 0.00 | 0.11 | 0.07 | 0.50 |
| 0.5 | 0.5 | 0 | 0.06 | 0.35 | 0.34 | 0.88 | 0.01 | 0.12 | 0.13 | 0.59 |
| 0.5 | 0.1 | 0.3 | 0.02 | 0.22 | 0.04 | 0.40 | 0.00 | 0.15 | 0.04 | 0.37 |
| 0.5 | 0.3 | 0.3 | 0.02 | 0.23 | 0.08 | 0.50 | 0.01 | 0.15 | 0.05 | 0.41 |
| 0.5 | 0.5 | 0.3 | 0.05 | 0.27 | 0.13 | 0.62 | 0.01 | 0.15 | 0.07 | 0.43 |
| 0.9 | 0.1 | 0 | 0.01 | 0.20 | 0.03 | 0.41 | 0.00 | 0.10 | 0.04 | 0.39 |
| 0.9 | 0.3 | 0 | 0.01 | 0.22 | 0.09 | 0.57 | 0.00 | 0.10 | 0.06 | 0.46 |
| 0.9 | 0.5 | 0 | 0.05 | 0.33 | 0.25 | 0.78 | 0.01 | 0.11 | 0.10 | 0.54 |
| 0.9 | 0.1 | 0.3 | 0.01 | 0.21 | 0.03 | 0.35 | 0.00 | 0.12 | 0.02 | 0.32 |
| 0.9 | 0.3 | 0.3 | 0.02 | 0.22 | 0.05 | 0.42 | 0.00 | 0.12 | 0.03 | 0.34 |
| 0.9 | 0.5 | 0.3 | 0.04 | 0.26 | 0.07 | 0.45 | 0.01 | 0.13 | 0.04 | 0.35 |
| Median | | | 0.02 | 0.23 | 0.08 | 0.54 | 0.01 | 0.13 | 0.06 | 0.44 |

*Notes*: For 18 models, non-inertial ($\rho_1 = 0$) and inertial ($\rho_1 = 0.8$) versions of Rule 1 and non-inertial ($\rho_2 = 0$) and inertial ($\rho_2 = 0.8$) versions of Rule 2 are considered. For each combination of model and rule, 5,000 samples of 100 observations of data are generated. A term structure regression like Eq. (17) is estimated for each sample, and the resulting distributions of $R^2$ values are summarized by $R^2_{\mathrm{L}}$ and $R^2_{\mathrm{U}}$, the 5 percent lower and upper critical values. The median values of these critical values across the 18 models are also reported.

## 6. The illusion of monetary policy inertia

The large estimated lag coefficients in the empirical partial adjustment policy rules appear to provide strong evidence of monetary policy inertia. However, such quarterly inertia is inconsistent with the very low interest rate forecastability in the term structure of interest rates. This section shows how the partial adjustment evidence in the empirical rules may be explained by a rationale other than policy inertia.

As a first step, note that there is a large literature that argues that partial adjustment models such as Rules 1 and 2 are difficult to identify and estimate empirically in the presence of serially correlated shocks (e.g., Griliches, 1967; Blinder, 1986; Hall and Rosanna, 1991; McManus et al., 1994). In particular, a standard policy rule with slow partial adjustment and no serial correlation in the errors will be difficult to distinguish empirically from a policy rule that has immediate policy adjustment but highly serially correlated shocks. Using the 1987–1999 data sample from Section 2, this latter form of Rule 1 is estimated as[28]

$$\begin{aligned} i_t = \underset{(0.24)}{1.24\bar{\pi}_t} + \underset{(0.10)}{0.33y_t} + \xi_t, \quad \xi_t = \underset{(0.06)}{0.92\xi_{t-1}} + \omega_t, \\ \sigma_\omega = 0.36, \quad \bar{R}^2 = 0.96. \end{aligned} \tag{18}$$

This rule assumes immediate adjustment ($\rho_1 = 0$) but allows for first-order serial correlation of the shocks with an AR(1) coefficient denoted $\rho_1^e = 0.92$. The corresponding estimated serially correlated shock version of Rule 2 is

$$\begin{aligned} i_t = \underset{(0.66)}{2.00\mathrm{E}_{t-1}\bar{\pi}_{t+4}} + \underset{(0.24)}{0.39\mathrm{E}_{t-1}y_t} + \xi_t, \quad \xi_t = \underset{(0.11)}{0.77\xi_{t-1}} + \omega_t, \\ \sigma_\omega = 0.68, \quad \bar{R}^2 = 0.87, \end{aligned} \tag{19}$$

with $\rho_2^e = 0.77$. These two estimated autocorrelated shock versions of Rules 1 and 2 display a fit to the data as well as estimates of $g_\pi$ and $g_y$ that are broadly comparable to the partial adjustment forms in Eqs. (3) and (6).

For a more rigorous comparison, the partial adjustment and serially correlated shocks rules can be nested in a single equation and tested directly (as in Hendry and Mizon, 1978). For Rule 1, this general nesting form is

$$i_t = \rho_1^a i_{t-1} + g_\pi \bar{\pi}_t + g_y y_t - \rho_1^b (g_\pi \bar{\pi}_{t-1} + g_y y_{t-1}) + \omega_t. \tag{20}$$

The hypothesis that policy Rule 1 is non-inertial but has serially correlated shocks is $\mathrm{H1_{SC}}$: $\rho_1^a = \rho_1^b \equiv \rho_1^e$. With this "common factor" restriction, Eq. (20) is the quasi-differenced form that matches the AR(1) shock rule, which is estimated above as Eq. (18). Alternatively, the hypothesis that the central bank follows a partial adjustment Rule 1 is $\mathrm{H1_{PA}}$: $\rho_1^b = 0$ (with $\rho_1^a \equiv \rho_1 \neq 0$). With this restriction, the estimated version of Eq. (20) matches the partial adjustment form (2).[29]

Unfortunately, it is difficult to obtain decisive direct empirical evidence against either of these hypotheses. Over the 1987–1999 sample, the $p$-value of the serially correlated shock hypothesis $\mathrm{H1_{SC}}$ is 0.18, while the $p$-value for the partial adjustment hypothesis $\mathrm{H1_{PA}}$ is 0.14. That is, over this sample, there is little evidence to reject either of these two forms. Even worse, the evidence appears quite fragile to even modest changes in the sample. For example, as shown in Table 5, in a slightly shorter sample, the serially correlated shock Rule 1 is rejected, while in a slightly longer sample, the partial adjustment Rule 1 is rejected. (The $p$-value of $\mathrm{H1_{SC}}$ is zero in the 1987–1996 sample, and the $p$-value of $\mathrm{H1_{PA}}$ is zero in the 1983–1999 sample.) Similarly fragile results are given in Table 5 for Rule 2 as well.

This difficulty in distinguishing partial adjustment from serially correlated shocks is consistent with the inventory adjustment econometrics literature cited above. The choice between these two forms of modeling dynamics depends crucially on

[28] Rule 1 with an AR(1) error is estimated via maximum likelihood, while Rule 2 with an AR(1) error is estimated with an instrumental variables version of the Hildreth–Lu procedure.

[29] Of course, the $\hat{g}_\pi$ and $\hat{g}_y$ in Eq. (20) would equal $(1 - \hat{\rho}_1)$ times the corresponding estimates in Eq. (2).

Table 5
Partial adjustment and serially correlated shock rule $p$-values

| Sample | | Rule 1 | | Rule 2 | |
|---|---|---|---|---|---|
| Start date | End date | $H1_{PA}$ | $H1_{SC}$ | $H2_{PA}$ | $H2_{SC}$ |
| 1987:4 | 1999:4 | 0.18 | 0.14 | 0.02 | 0.18 |
| 1987:4 | 1996:4 | 0.27 | 0.00 | 0.56 | 0.11 |
| 1983:4 | 1999:4 | 0.00 | 0.06 | 0.00 | 0.26 |

*Note*: The table entries are $p$-values of the partial adjustment (PA) or serially correlated shocks (SC) restricted versions of Rules 1 and 2.

separating the influences of contemporaneous and lagged regressors, which are especially difficult to untangle for empirical monetary policy rules for several reasons. First, the arguments of the rules—four-quarter inflation and the output gap—are highly serially correlated, so distinguishing the effect of, say, $\bar{\pi}_t$ from $\bar{\pi}_{t-1}$ is not easy. Second, the arguments of the rules are not exogenous (as is often assumed in the inventory adjustment literature) but depend crucially on past interest rates. Third, only short data samples of plausibly consistent rule behavior are available with a limited amount of business cycle variation in output and inflation. Fourth, there is some uncertainty about the appropriate arguments of the historical policy rule. Finally, the actual interest rates are set on the basis of real-time data on output and inflation, which also makes it difficult to determine the correct dynamics (see Rudebusch, 1998, 2001, 2002, and the discussion below). Indeed, the near-observational equivalence of partial adjustment and serially correlated shocks for monetary policy rules provides a key motivation for examining the indirect term structure evidence as above.

The estimated partial adjustment policy rules failed the indirect term structure test in Section 5 by implying too much interest rate forecastability, but the serially correlated shocks in the near-observational-equivalent estimated rules in Eqs. (18) and (19) may also translate into interest rate forecastability. Certainly, in the general form of Eq. (20), as $\rho_1^a$ increases for fixed $\rho_1^b$, the forecastability of interest rates should increase as it did for the standard partial adjustment model. Indeed, in Fig. 1, with $\rho_1^b$ equal to zero, forecastability increased with $\rho_1^a \equiv \rho_1$. However, in the general case, this intuition ignores the offsetting effect on forecastability of simultaneously increasing $\rho_1^b$. Since $\bar{\pi}_t$ and $y_t$ are persistent processes, as $\rho_1^b$ increases, the term $g_\pi \bar{\pi}_t + g_y y_t - \rho_1^b(g_\pi \bar{\pi}_{t-1} + g_y y_{t-1})$ becomes less predictable. This effect is illustrated in Fig. 2, which examines the forecastability of interest rates (again as measured by the $R^2$ of the term structure regression) with the rule in Eq. (20) (in a model with $\mu_\pi = 0.3$, $\mu_r = 0.5$, and $\mu_y = 0$). As shown by the downward-sloping thin solid line, with $\rho_1^a$ set equal to 0.7, as $\rho_1^b$ increases from zero to 0.6, the forecastability of $\Delta i_{t+2}$ declines. The thin dashed line gives a similar result for Rule 2. The thicker lines in Fig. 2 give the effect on interest rate predictability of *simultaneously* increasing $\rho_1^a$ and $\rho_1^b$. In particular, the thick solid line shows the forecastability of $\Delta i_{t+2}$ for the rule with $\rho_1^a = \rho_1^b \equiv \rho_1^e$, which matches the AR(1) shock rule. As $\rho_1^e$, the persistence

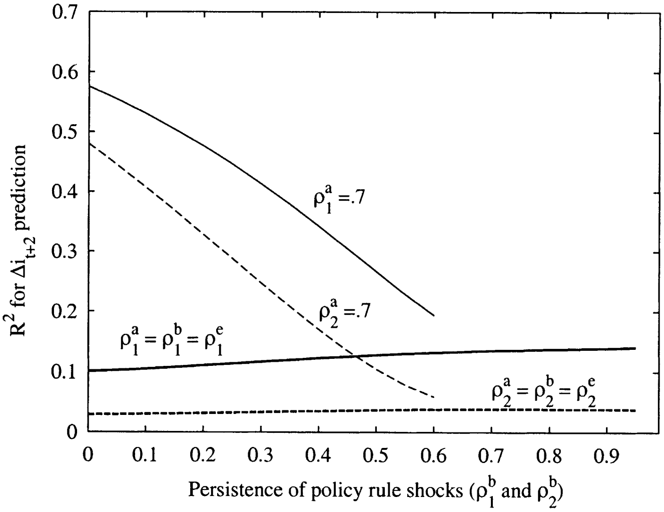


Fig. 2. Interest rate forecastability and policy shock persistence. The solid lines show the effects on interest rate predictability of varying $\rho_1^b$ in the policy rule $i_t = \rho_1^a i_{t-1} + g_\pi \bar{\pi}_t + g_y y_t - \rho_1^b(g_\pi \bar{\pi}_{t-1} + g_y y_{t-1}) + \omega_t$. In the upper solid line, $\rho_1^a$ is constant. In the lower line, $\rho_1^a = \rho_1^b$, so the rule matches a serially correlated shocks version of Rule 1. The dashed lines show similar results for Rule 2.

Table 6
Serially correlated shocks and interest rate forecastability

| AR(1) shock | $\Delta i_{t+2}$ forecastability | | $\Delta i_{t+3}$ forecastability | |
|---|---|---|---|---|
| policy rule | $R^2_{\mathrm{L}}$ | $R^2_{\mathrm{U}}$ | $R^2_{\mathrm{L}}$ | $R^2_{\mathrm{U}}$ |
| Rule 1 ($\rho_1^e = 0.90$) | 0.05 | 0.31 | 0.04 | 0.28 |
| Rule 2 ($\rho_2^e = 0.90$) | 0.02 | 0.16 | 0.01 | 0.14 |

*Notes*: For 18 models, a serially correlated shocks version of Rules 1 and 2 is considered (with $\rho_1^e = \rho_2^e = 0.90$). For each combination of model and rule, 5,000 samples of 100 observations of data are generated. Term structure regressions like Eqs. (16) and (17) are estimated for each sample, and the resulting distributions of $R^2$ values are summarized by $R^2_{\mathrm{L}}$ and $R^2_{\mathrm{U}}$, the 5 percent lower and upper critical values. Only the median values of these critical values across the 18 models are reported.

of the policy rule shocks, increases, the forecastability of interest rate changes is remarkably unaffected. The thick dashed line shows a similar result for Rule 2.

Fig. 2 gives analytical, asymptotic results, so Table 6 provides some relevant small sample evidence. Rules 1 and 2 take forms similar to the ones above, with $g_\pi = 1.5$, $g_y = 0.8$, $\rho_1 = \rho_2 = 0$, and an AR(1) shock calibrated with $\rho_1^e = \rho_2^e = 0.90$ and $\sigma_\omega = 0.4$. The resulting $R^2_{\mathrm{L}}$ and $R^2_{\mathrm{U}}$ pairs for the $\Delta i_{t+2}$ and $\Delta i_{t+3}$ prediction regressions show that these rules with serially correlated shocks display little interest rate forecastability, which is consistent with the historical results. (For brevity, only the median confidence intervals across the 18 models are provided.)

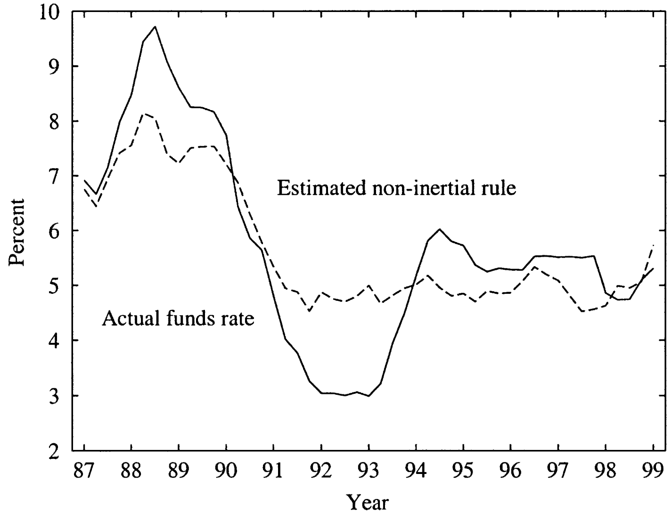


Fig. 3. Actual funds rate and an estimated non-inertial rule. The solid line is the historical funds rate. The dashed line shows the fitted values from an estimated non-inertial version of Rule 1: $i_t = 1.24\bar{\pi}_t + 0.33y_t$.

Much has been written about monetary policy "shocks", such as the $\xi_t$, in the literature, so it is useful to provide some economic interpretation of these persistent rule deviations. Recall the original analysis of Taylor (1993), which put forward a description of monetary policy that did not involve interest rate smoothing or partial adjustment. Taylor argued that recent historical monetary policy had followed a rule only as a guide, so deviations from the rule during various episodes were an appropriate response to special circumstances, not evidence of partial adjustment. This view is illustrated in Fig. 3, which displays the historical values of the funds rate (solid line) and the fitted values (dashed line) from the estimated non-inertial Rule 1 in Eq. (18), which allows for serially correlated shocks. The associated large persistent shocks, that is, the deviations between the two lines, appear to correspond to several special episodes.[30] Most notably, the deviations in 1992 and 1993 are commonly interpreted as responses to a disruption in the flow of credit. As Fed Chairman Alan Greenspan testified to Congress on June 22, 1994: "Households and businesses became much more reluctant to borrow and spend and lenders to extend credit—a phenomenon often referred to as the "credit crunch". In an endeavor to defuse these financial strains, we moved short-term rates lower in a long series of steps that ended in the late summer of 1992, and we held them at unusually low levels through the end of 1993—both absolutely and, importantly, relative to inflation". Thus, this episode appears better described as a persistent "credit crunch" shock than as a sluggish partial adjustment to a known desired rate. Similarly, a worldwide

[30] Kozicki (1999, p. 24) also makes the point that "... information and events outside the scope of Taylor-type rule specifications..." often appear to influence policy actions.

financial crisis appeared to play a large role in lowering rates in 1998 and 1999, and commodity price scares pushed rates up in 1988–1989 and 1994–1995.[31]

This description of credit crunches and financial crises clarifies the fact that these rule deviations are not "exogenous policy shocks", that is, actions undertaken by central bankers that are independent of the economy (and the focus of the VAR literature). Instead, these deviations are endogenous responses to a variety of influences that cannot be captured by some easily observable variable such as output or inflation.[32] In terms of the Taylor rule in Eq. (1), one interpretation of the rule deviations is that they are fluctuations in the equilibrium real rate, $r^*$. For example, as suggested by Greenspan's testimony, a disruption of credit supply could be treated as a temporary lowering of the equilibrium real rate, and in response, the Fed lowers the funds rate relative to readings on output and inflation. It should be clear that the modeling of these shocks to the Taylor rule as an AR(1) process, as in Eq. (18), is simply a convenient econometric approximation.[33]

A complementary rationale for the serially correlated shocks has also been discussed in the policy rules literature. Several recent papers (e.g., Smets, 1999; Rudebusch, 2001, 2002; Orphanides et al., 1999) have stressed that setting monetary policy according to a Taylor rule requires relying on a "real-time" estimate of the output gap. They advocate a "real-time analysis" (as defined by Diebold and Rudebusch, 1991), which uses the sequential information sets that were actually available as history unfolded. The available historical data suggest that the real-time output gap estimates, denoted $y_{t|t}$, are very noisy versions of the final estimates, $y_t$. The large and persistent revisions, $n_t$, can be defined by $y_{t|t} = y_t + n_t$. In this case, even if the central bank follows Rule 1 with no partial adjustment or error in real time,

$$i_t = g_\pi \bar{\pi}_t + g_y y_{t|t}, \tag{21}$$

the econometrician working with the final data will estimate

$$i_t = \hat{g}_\pi \bar{\pi}_t + \hat{g}_y y_t + k_t, \tag{22}$$

where the error $k_t = \hat{g}_y n_t$ is the highly serially correlated real-time data noise. Lansing (2002) provides a careful simulation study that demonstrates the potential

[31] Federal Reserve Governor Laurence Meyer (1999, p. 7) had this explanation for the easing of policy during late 1998: "There are three developments, each of which, I believe, contributed to this decline in the funds rate relative to Taylor Rule prescription. The first event was the dramatic financial market turbulence, following the Russian default and devaluation. The decline in the federal funds rate was, in my view, appropriate to offset the sharp deterioration in financial market conditions, including wider private risk spreads, evidence of tighter underwriting and loan terms at banks, and sharply reduced liquidity in financial markets."

[32] Rules 1 and 2 may appear too parsimonious so that the persistent deviations reflect a serially correlated omitted variable; however, as noted above, the empirical reaction function literature, including monetary VARs, has placed the proverbial kitchen sink on the right-hand side in attempts to explain the policy rate, yet serially correlated errors remain, which are modeled through lagged interest rates and partial adjustment. Again, see Rudebusch (1998) and Goodfriend (2000).

[33] As an alternative, Gerlach and Schnabel (2000) include in their estimated Taylor rule for Europe a dummy variable intercept shift for a large persistent rule deviation. They find (p. 167) that a European Taylor rule fits well without partial adjustment but with "… dummies for the period 1992:3–1993:3 to control for policy responses to intra-European exchange market pressures in this period".

effectiveness of such real-time output gap errors to account for the spurious evidence of policy inertia in exactly this fashion.[34] Indeed, based on a reconstruction of real-time output gap data for the U.S., Mehra (2002) reports that the evidence for policy inertia and interest rate smoothing is reduced when estimating partial adjustment rules using the real-time data. Real-time output gap revisions may not be a complete explanation because there are estimated reaction functions with significant inertia that do not include an output gap (for example, McNees, 1992; McCallum and Nelson, 1999; Fair, 2000; and the VAR interest rate equations); however, it seems likely that in general the real-time information set is an important element in accounting for spuriously inertial estimated policy rules.

## 7. Conclusion

Empirical monetary policy rules with large estimated coefficients on the lagged policy interest rate, which are very prevalent in the literature, are widely interpreted as indicating a sluggish adjustment of the policy rate to its determinants—on the order of only about 20 percent per quarter. This partial adjustment implies predictable future changes in the policy rate over horizons of several quarters, which does not accord with the lack of information about such changes in financial markets. This paper proposes a resolution of this empirical inconsistency by providing an alternative interpretation of the large lag coefficients in the estimated policy rules. These coefficients reflect serially correlated or persistent special factors or shocks that cause the central bank to deviate from the policy rule.

This argument uses indirect term structure evidence to dismiss the interest rate smoothing interpretation of the partial adjustment rule. As noted above, it appears difficult to develop direct evidence against the partial adjustment rule (in the form of non-rejection of the $\rho = 0$ hypothesis). In particular, the uncertainty in modeling the desired policy rate (given the endogeneity of its determinants, the real-time nature of the information set, as well as the small samples available) makes any direct evidence from estimated rules fragile. For example, the rule with partial adjustment and the rule with serially correlated shocks both appear to fit the data as empirical reaction functions. However, they have very different economic interpretations. In the former rule, persistent deviations from an output and inflation response occur because policymakers are slow to react. In the latter rule, these deviations reflect the policymaker's response to other persistent influences. The two types of rules can be distinguished, however, by their very different implications for the term structure. Only the serially correlated shocks rule is consistent with the historical evidence showing that the term structure is largely uninformative about the future course of the policy rate.

There may be other possible reconciliations of the policy rule and term structure empirical results. For example, it may be that the rational expectations hypothesis of

[34] Much earlier, Goodfriend (1985) also showed that measurement error in the variables determining money demand could result in spuriously significant partial adjustment lags.

the term structure cannot be applied and the associated term structure interpretations above are spurious. One way in which this hypothesis may fail is that expectations are not rational, but this would undermine many aspects of any explicitly forward-looking macroeconomic modeling exercise such as the one above. Or term premia for short-term interest rates may be even more volatile than assumed above; however, if rates are driven by volatile term premia, then it seems unlikely that they can communicate the subtle expectations of future monetary policy as required in the literature on optimal monetary policy inertia.

It is also possible that there is some intermediate case of partial adjustment, a $\rho_1$ or $\rho_2$ of 0.4, say, along with some serially correlated shocks, that is not strictly rejected by the term structure evidence. However, it should be noted that while real-world discussions of monetary policy sometimes mention the "incrementalism" and "gradualism" of smoothing the policy rate over the next several weeks, there is no acknowledgment of *quarterly* interest rate smoothing.[35] As the *New York Times* (July 26, 2000) summarized of recent Congressional testimony: "Alan Greenspan, the Federal Reserve chairman, said today that the central bank's decision about whether to raise interest rates again at its meeting next month would hinge in large part on economic data released in coming weeks." That is, there was little if any pent-up pressure from the past for further adjustment.

In future research, the empirical rules given in Section 6 can be improved as further effort is made in estimating rules without the crutch of partial adjustment. Given the similar estimates above of $g_\pi$ and $g_y$ across rules, it may be that past conclusions about these coefficients, as in Clarida et al. (2000), are robust to the exact formulation of serial correlation in the rule. However, the lagged policy rate, though useful in mopping up residual serial correlation, should not be given a structural partial adjustment interpretation with regard to central bank behavior. In particular, using the partial adjustment rule in a model as a representation of historical policy (as in Levin et al., 1999, and many other studies) may give misleading results, especially about the nature of optimal policy inertia.

With regard to optimality, the maintained hypothesis of economics for central banks, as for other agents in the economy, is that the non-inertial policy rule apparently used in practice is optimal, and certainly, the rule can be rationalized as such in particular models as in Table 2. However, it should be stressed that there are many aspects of the monetary policy process still to be modeled, especially imperfect credibility and uncertainty (see Rudebusch, 2001).

Also, the absence of partial adjustment does not mean that central banks are not trying to influence long-term interest rates. However, in order to influence the long rate, central banks only must present a clear path for the policy rate that can shape expected future rates. The partial adjustment rule provides one such path, but it is not the only one. As noted by Goodfriend (1991) and Rudebusch (1995), an ex ante constant path, which is approximately what the non-inertial rules deliver, is another obvious choice.

[35] And as noted in the introduction, models of monthly interest rate smoothing imply very little if any quarterly interest rate smoothing.

Finally, further careful analysis of the empirical policy rule is required in modeling and identifying the shocks. Section 6 provides a simple formulation for adding shocks to a policy rule. A better specification may link persistent shocks in both the rule and the rest of the model. A bout of credit frictions or impediments may lower the equilibrium real rate and provide a persistent negative shock to the policy rule and to the output equation as well (see Rudebusch, 2001). Alternatively, an idiosyncratic inflation scare may provide a shock to the rule and to inflation expectations more broadly.

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